NO TREND IN THE INTERGENERATIONAL TRANSMISSION OF DIVORCE*

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Previous studies on trends in the intergenerational transmission of divorce have produced mixed findings, with two studies (McLanahan and Bumpass 1988; Teachman 2002) reporting no trend in divorce transmission and one study (Wolfinger 1999) finding that divorce transmission has weakened substantially. Using a stratified Cox proportional hazard model, we analyze data from the National Survey of Families and Households and find no evidence for any trend in divorce transmission. To reconcile apparent differences in results, we note that the General Social Survey data used by Wolfinger lack information on marital duration, permitting analysis only for whether respondents have divorced by interview. As a result, an apparent decline in divorce transmission could be due to inadequate adjustments for the longer exposures to risk by earlier marriage cohorts, yielding a higher probability of divorce by interview for earlier cohorts relative to more recent cohorts even if divorce risks are identical across all marriage cohorts. We confirm this possibility by using a series of discrete-time hazard logistic regressions to investigate the sensitivity of estimates of trends in divorce transmission to different adjustments for exposure to risk. We conclude that there has been no trend in the intergenerational transmission of divorce.

The intergenerational transmission of divorce has been well documented. Children of divorce are more likely to divorce than those who grew up in two-parent families (see, e.g., Amato 1996; Bumpass, Castro Martin, and Sweet 1991; Kiernan and Cherlin 1999). There is, however, less consensus concerning trends in the intergenerational transmission of divorce. McLanahan and Bumpass (1988), analyzing data from the 1982 National Survey of Family Growth, found that the intergenerational transmission of divorce risks for those married after 1970 was similar for all marriage cohorts in their sample, suggesting no trend by marriage cohort in divorce transmission. Teachman (2002), pooling data from five waves of the National Survey of Family Growth to investigate differences in divorce rates for a number of observed characteristics of respondents, found that the effect of parental divorce remained fairly constant for marriages contracted between 1950 and 1984, a finding again suggesting no trend by marriage cohort in the intergenerational transmission of divorce. By contrast, Wolfinger (1999), pooling data from the 1973–1996 General Social Survey (GSS), concluded that divorce transmission declined by almost 50% between 1973 and 1996.

In this article, we address these apparent discrepancies concerning trends in the intergenerational transmission of divorce. Because the GSS lacks information on marital duration, Wolfinger analyzed trends in whether a respondent had ever divorced by GSS interview. However, if divorce rates are identical across marriage cohorts, more respondents in earlier marriage cohorts would be observed to divorce relative to respondents in later marriage cohorts simply by virtue of longer exposures to the risk of divorce. If Wolfinger's controls for exposure to risk are inadequate, his conclusion that divorce transmission has declined could be a methodological artifact.[1]

1. By contrast, both McLanahan and Bumpass (1988) and Teachman (2002) analyzed divorce risks by using hazard regression methods that adequately controlled for exposure to the risk of divorce.

To investigate this possibility, we report three sets of analyses using data from the 1987–1988 National Survey of Families and Households (NSFH). We show that (1) modeling the probability of divorce via logistic regression yields Wolfinger's finding of a declining trend of divorce transmission, but (2) modeling divorce risks using a Cox proportional hazard model suggests no trend in divorce transmission. We then use discrete-time hazard logistic regressions to show how estimates of divorce transmission vary with how accurately we model differential exposure to risk. Our results yield increasingly small estimates for the decline in divorce transmission as we better approximate a continuous-time Cox model using increasingly small intervals to better model differential exposure to risk within a discrete-time hazard framework. Conversely, estimates for divorce transmission increase in magnitude as we increasingly approximate a logistic regression using larger intervals that model differential exposure to risk less well. These results lead us to conclude that there is no trend in the association between a parental divorce and offspring's own *risk* of divorce.

DATA AND METHODS

Data

We use data from the 1987–1988 National Survey of Families and Households (NSFH). The NSFH consists of a national probability sample of people ages 19 and over who resided in the United States in 1987 and 1988, with a main sample of 9,643 cases and an oversample of 3,374 cases of minorities, single-parent families, families with stepchildren, cohabiting couples, and recently married persons (Sweet, Bumpass, and Call 1988). The NSFH contains detailed retrospective marital histories—unlike the GSS, which contains data on a respondent's age at first marriage and whether the respondent had divorced by GSS survey but which lacks information on how and when a marriage ended. Thus, the marriage histories in the NSFH allow us to contrast results from models of the probability of divorce by NSFH survey and models of divorce rates for successive marriage cohorts. We examine the probability of divorce by NSFH survey using a dichotomous measure of whether the respondent reported having ended his/her first marriage through separation or divorce. To examine divorce rates, we use a dichotomous indicator of whether the respondent's first marriage ended in separation or divorce and a continuous measure of marital duration, measured to the nearest month, censoring first marriages at the time of spousal death or interview.

Parental divorce is constructed using the retrospective history of childhood family structure in the NSFH, in which respondents reported family structure (including parental divorce) through age 18. Marriage cohort is constructed by the calendar year in which the respondent reported being married the first time. Following Wolfinger (1999) and Teachman (2002), we use a simple linear interaction of parental divorce and time (survey date in Wolfinger's analyses, marriage cohort in our analyses and Teachman's) to assess trends in the intergenerational transmission of divorce.[2]

We include the following control variables in our Cox, logistic, and discrete-time hazard regressions: the respondent's sex, race (black vs. nonblack), education, age at first marriage, and religion (Catholic vs. non-Catholic); the interaction of religion and parental divorce; urban/rural residence; whether the respondent is an only child; respondent's occupational prestige; and parental education. We restrict our analytic sample to ever-married respondents, omitting respondents whose marital status was unknown and respondents with illogical marital histories or with missing data on how and when a first marriage ended. We also restrict the sample to respondents under 90 years of age and omit respondents with

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2. We also examine alternative nonlinear, piecewise-spline specifications of this interaction (results not reported), with our results robust across these alternative specifications for trends.

Table 1. **Descriptive Statistics, Means and Standard Deviations, NSFH 1987–1988**

Variables	Mean	SD
Dependent Variables		
Ever divorced by survey interview	.27	
Marriage duration (in months)	234.07	190.23
Independent Variables		
Parents divorced	.09	
Marriage cohort (calendar year – 1900)	61.81	16.40
Male	.45	
Black	.09	
Catholic	.26	
Education		
Less than high school	.23	
Some college	.19	
Bachelor's degree	.12	
Postgraduate education	.07	
Occupational prestige	38.00	14.92
Occupational prestige missing	.39	
Only child	.06	
Residence in rural area	.27	
Age at first marriage	22.61	4.77
Parental education		
Less than high school	.42	
Some college	.07	
Bachelor's degree	.05	
Postgraduate degree	.02	
Parental education missing	.13	
N		10,216

Notes: Means and standard deviations are weighted. Sample size *N* is unweighted.

missing data on date of birth, religion, and number of siblings, as well as respondents who did not speak English. We also include dummy variables for missing data on age at first marriage, parental education, and respondent's occupational prestige. The final analytic sample includes 10,216 respondents, thus retaining about 78% of the full NSFH sample. Descriptive statistics are in Table 1.

Statistical Models

To model trends in the intergenerational transmission of the probability of divorce, we estimate a logistic regression of the following form:

$$\log\left(\frac{p}{1-p}\right) = \beta_0 + \beta_1 PaDiv + \beta_2 MarCoh + \beta_3(PaDiv \times MarCoh) + \boldsymbol{\beta}\mathbf{x}, \tag{1}$$

where p denotes the probability of ever divorcing by the time of the NSFH interview, *PaDiv* denotes a dummy variable equal to 1 if the respondent's parents divorced, *MarCoh* denotes marriage cohort (in calendar year minus 1900), and **x** denotes a vector of control variables. Under this specification, a negative value of β_3 corresponds to a declining trend in intergenerational transmission for the probability of divorce by survey. However, because this logistic regression model does not adjust for exposure to risk, earlier marriage cohorts will have longer exposures to the risk of divorce than later marriage cohorts.[3] It is thus possible for divorce risks to be identical for successive marriage cohorts, but for more of those from earlier cohorts to have divorced by date of survey simply by virtue of differential exposure to risk.

To model trends in the intergenerational transmission of the duration-specific risk of divorce, we estimate a Cox proportional hazard model stratified by three groups of marriage cohorts, 1935–1954, 1955–1974, and 1975–1984[4]:

$$r_s(t) = r_{0s}(t) \cdot \exp\left(\beta_1 PaDiv + \beta_2 MarCoh + \beta_3 (PaDiv \times MarCoh) + \beta \mathbf{x}\right), \tag{2}$$

where $r_s(t)$ denotes the risk of divorce in stratum s at marital duration t, and $r_{s0}(t)$, denoting the baseline hazard in stratum s, is an unspecified nonnegative function specific to each stratum s, with the β's constrained to be equal across strata.

To examine how estimates of divorce transmission vary with the accuracy of adjustments for differential exposure to risk, we specify a series of discrete-time hazard logistic regressions. In most applications, analysts using discrete-time hazard methods choose a single value for the discrete time period Δt, constructing, for example, a file of person-year or person-month records (Allison 1982). By contrast, we construct 10 person-period files in which we vary Δt and the corresponding number of discrete time periods T. These 10 files range from a file with an infinite interval of exposure ($\Delta t = \infty$) and only one discrete time period per person ($T = 1$) to a file with a one-year interval ($\Delta t = 1$) and a maximum of 70 discrete time periods per person ($T = 70$). We then estimate 10 discrete-time hazard logistic regressions of the following form:

$$\log\left(\frac{q_t}{1-q_t}\right) = \sum_{t=1}^{T} \alpha_t D_t + \beta_1 PaDiv + \beta_2 Marcoh + \beta_3 (PaDiv \times Marcoh) + \beta \mathbf{x}, \tag{3}$$

where q_t is the predicted probability of divorce for the marital duration corresponding to the discrete time period t conditional on a surviving marriage in period $t-1$, and D_t is a set of dummy variables indicating period t. These models increasingly approximate a continuous-time proportional hazard model as Δt shrinks and the number of periods increases, with $\sum_{t=1}^{T} \alpha_t D_t$ and the logistic estimates of the β's in Eq. (3) converging to their continuous-time hazard counterparts (Allison 1982). At the other extreme, when $\Delta t = \infty$, there is only one period per person ($T = 1$); $\sum_{t=1}^{T} \alpha_t D_t$ in Eq. (3) then reduces to a single intercept, β_0, thus yielding a logistic regression identical to the logistic regression in Eq. (1).

RESULTS

Table 2 presents results for trends in the transmission of the probability of divorce by survey using logistic regressions and for trends in the transmission of divorce risks using a

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3. Wolfinger (1999) employed a somewhat different specification from Eq. (1), replacing year of marriage cohort by year of GSS survey and by including a control for the difference between age at GSS survey and age at first marriage, with this latter variable specified as a piecewise linear spline.

4. We stratify by marriage cohort because of evidence that the proportionality assumption is violated across these strata (see Li and Wu 2006).

Table 2. **Trends in Probabilities of Divorce Transmission (logistic regressions) Versus Trends in Rates of Divorce Transmission (stratified Cox proportional hazard models)**

	Logit Model, Divorce Probabilities		Stratified Cox Model, Divorce Rates	
	Without Interaction	With Interaction	Without Interaction	With Interaction
Parental Divorce	0.45**	1.27**	0.41**	0.45
	(0.09)	(0.38)	(0.07)	(0.34)
Marriage Cohort	–0.003	–0.002	0.030**	0.031**
	(0.002)	(0.002)	(0.005)	(0.005)
Parental Divorce × Marriage Cohort		–0.012*		–0.001
		(0.005)		(0.005)
Log-Likelihood	–5,759.58	–5,756.74	–21,176.74	–21,176.73

Notes: Robust standard errors are in parentheses. All models are weighted and include control variables of respondent's sex, race, education, and religion; the interaction of religion and parental divorce; urban/rural residence; whether s/he is the only child; occupational prestige; and his/her parent's education. Sample size $N = 10,216$.

$^*p < .05$; $^{**}p < .01$

stratified Cox proportional hazard model. The logistic regression results appear to indicate a significant decline in the intergenerational transmission of the probability of divorce by date of survey ($\beta = -0.012$, $z = -2.30$), with the relative odds declining by about 11% for marriages contracted 10 years apart ($[1 - \exp(-0.012 \times 10)] \times 100\% = 11\%$). By contrast, the Cox regression results indicate a trivial and statistically insignificant trend in the transmission of divorce rates ($\beta = -0.0006$, $SE = 0.0050$, $z = -0.12$). A comparison of the likelihoods for two Cox models, one including the interaction of parental divorce and marriage cohort and the other excluding this term, shows a negligible increment in fit when including this interaction. These results are consistent with the argument that without adequate control for exposure to the risk of divorce, more of those from earlier marriage cohorts will have divorced by the date of survey simply by virtue of differential exposure to risk, with no corresponding trend in the intergenerational transmission of divorce risks by marriage cohort.

Table 2 shows that including the interaction between parental divorce and marriage cohort increases the standard error for parental divorce substantially (from 0.09 to 0.38 in the logit models and from 0.07 to 0.34 in the Cox models). This raises the possibility of collinearity between the interaction term and one or both of the main effect terms for parental divorce or marriage cohort. Note also that the estimated logistic regression coefficients for marriage cohort are small, negative, and not statistically significant ($\beta = -0.003$, without interaction; $\beta = -0.002$, with interaction), indicating that later marriage cohorts have a similar (and possibly smaller) probability of divorce relative to earlier marriage cohorts. By contrast, the estimated Cox regression coefficients for marriage cohort are positive and statistically significant ($\beta = 0.030$ without the interaction term; $\beta = 0.031$ with the interaction), indicating that more recent marriage cohorts have higher divorce rates than earlier marriage cohorts. Thus, estimates from the Cox model are consistent with studies that have found that more recent marriage cohorts have higher divorce rates relative to earlier marriage cohorts (see, e.g., Preston and McDonald 1979), but estimates from the logistic regressions are inconsistent with this finding.

Sensitivity of Estimated Trends in Divorce Transmission to Adjustments for Exposure to Risk

In Table 3, we assess the sensitivity of estimated trends in divorce transmission to the accuracy with which we adjust exposure to risk by estimating a series of discrete-time hazard logistic regressions. As noted above, we specify a series of discrete-time logistic regressions replicating the logistic regression results in Table 2 at one extreme and approximating a continuous-time Cox model at the other extreme. Between these extremes, we examine results by varying the length of periods in our discrete-time hazard models, with the duration of each period and the resulting number of periods per person (and thus total number of person-period records in each of the 10 constructed files) reflecting the accuracy with which exposure to the risk of divorce is adjusted.

Results in the first column of Table 3 use a discrete-time hazard specification in which there is only one record per person and in which the single period is infinite in duration. This specification is identical to the logistic regression model in Eq. (1) and thus yields estimates identical to those reported in Table 2. But these estimates can now be seen within a discrete-time hazard framework as resting on the assumption that duration dependence in the risk of divorce can be modeled with only one parameter, a strong and implausible assumption. As we relax this assumption by increasing the number of duration intervals and by decreasing interval length in subsequent models, the coefficient for the interaction of parental divorce and marriage cohort falls in magnitude and statistical significance, with this coefficient not statistically significant in columns 4 to 10 (i.e., when Δt is 15 years or shorter, or when there are five or more periods per person). Note, in particular, that the lack of significance for the interaction coefficient of parental divorce and marriage cohort is not due to inflated standard errors; instead, the standard error for this coefficient is relatively stable across models and, indeed, reduces slightly with the number of duration intervals specified. Thus, these results suggest that the coefficient for the trend in the intergenerational transmission of divorce rates is highly sensitive to model assumptions about how the risk of divorce varies with marriage duration, with models that make weaker assumptions yielding coefficients that are small in magnitude and not statistically significant. These results provide additional evidence that inadequate control for exposure to the risk of divorce yields an artifactual decline in trends in divorce transmission.

Overall, our results provide strong evidence that there is no trend in the intergenerational transmission of divorce risks by marriage cohort. They show that a declining trend in the probability of divorce by survey date reflects the fact that more respondents in earlier cohorts will be observed to divorce by survey by virtue of longer exposure to risk.

DISCUSSION

Our results, like those of a number of other studies, provide evidence that children who experience the divorce of their parents are themselves more likely to divorce. But contrary to Wolfinger's (1999) claim that the intergenerational transmission of divorce has declined by nearly 50%, we find no trend in the intergenerational transmission of divorce risks, a result that mirrors the findings of McLanahan and Bumpass (1988) and Teachman (2002). We find that the discrepancy between these findings is due to differential exposure to the risk of divorce by successive marriage cohorts, with those married earlier having longer durations of exposure to divorce and those married later having shorter durations of exposure. Thus, caution is required when assessing trends in divorce transmission with data lacking information on marriage duration or when differences in exposure to risk are not adequately adjusted. By contrast, our analyses using marriage history data in the NSFH allow us to model duration-specific divorce rates using hazard regression methods, which automatically adjust for differences in exposure to risk. These data also let us assess how trend estimates vary under progressively more accurate adjustments for exposure to risk. Overall, these results

Table 3. Trends in the Intergenerational Transmission of Divorce: Discrete-Time Hazard Models Specified in Logistic Regression on Different Person-Period Files (with varying lengths of the discrete periods and varying maximum numbers of periods per person)

Length of Discrete Periods, in Years (t)	∞	35	25	15	10	5	4	3	2	1
Maximum Number of Periods per Person (T)	1	2	3	5	7	14	18	24	35	70
Parental Divorce	1.27**	1.22**	1.18**	0.99**	0.86**	0.72*	0.66*	0.66*	0.62*	0.57
	(0.38)	(0.35)	(0.33)	(0.30)	(0.30)	(0.30)	(0.30)	(0.30)	(0.30)	(0.31)
Marriage Cohort	−0.002	−0.001	0.003	0.011**	0.016**	0.022**	0.024**	0.025**	0.027**	0.029**
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Parental Divorce × Marriage Cohort	−0.012*	−0.012*	−0.011*	−0.008	−0.006	−0.004	−0.003	−0.004	−0.003	−0.002
	(0.005)	(0.005)	(0.005)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
N of Person-Period Records	10,216	11,912	13,515	18,225	22,991	39,784	48,352	62,621	90,979	171,139

Notes: Robust standard errors are in parentheses. All models are weighted and include controls of respondent's sex, race, education, and religion; the interaction of religion and parental divorce; urban/rural residence; whether s/he is the only child, occupational prestige; and his/her parent's education. The number of respondents is 10,216 across all models.

*$p < .05$; **$p < .01$

provide a cautionary tale that highlights the importance of appropriately controlling for differential exposure to risk when analyzing trends in demographic behaviors.

Whether the intergenerational transmission of divorce has declined over time or remained stable is not simply of methodological interest, but instead is a substantive issue that speaks to important and ongoing debates in the social sciences. Some have argued that as divorce has become more widespread, growing up in a divorced family has become less stigmatized, with the result that differences in child well-being (including offspring's marital stability) may have declined with time (Amato and Keith 1991; Wolfinger 1999). Although this stigmatization hypothesis is plausible, we conclude that it is inconsistent with observed trends in divorce transmission (see also Diekmann and Engelhardt 1999). More generally, growing up in a single-parent or stepparent family has been argued to be not merely associated with social and economic disadvantages, but plausibly considered as a cause of these disadvantages (Amato 2003; Cherlin 1999; McLanahan and Sandefur 1994). If childhood family structure is indeed a cause, the intergenerational transmission of family behaviors may be a central component in reproducing and maintaining inequality (Biblarz and Raftery 1999; McLanahan 1985), but others have presented evidence suggesting that the intergenerational transmission of poverty and of family structure follow independent pathways (Musick and Mare 2006). Whatever the mechanism underlying the intergenerational transmission of various behaviors, trends in phenomena such as divorce transmission are critical to such debates.

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